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Intermittent sea-level acceleration

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Abstract

Using instrumental observations from the Permanent Service for Mean Sea Level (PSMSL), we provide a new assessment of the global sea-level acceleration for the last ~ 2 centuries (1820–2010). Our results, obtained by a stack of tide gauge time series, confirm the existence of a global sea-level acceleration (GSLA) and, coherently with independent assessments so far, they point to a value close to 0.01 mm/yr^2 . However, differently from previous studies, we discuss how change points or abrupt inflections in individual sea-level time series have contributed to the GSLA. Our analysis, based on methods borrowed from econometrics, suggests the existence of two distinct driving mechanisms for the GSLA, both involving a minority of tide gauges globally. The first effectively implies a gradual increase in the rate of sea-level rise at individual tide gauges, while the second is manifest through a sequence of catastrophic variations of the sea-level trend. These occurred intermittently since the end of the 19th century and became more frequent during the last four decades.

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Keywords: Sea level rise, Sea level acceleration, Tide gauge observations

9 1. Introduction

10 In view of their impact on coastal hazard and society, the problems of
11 secular sea-level rise and of future sea-level trends are the subjects of ex-
12 tensive research (see e.g. Bindoff et al. 2007, Rahmstorf 2007, Cazenave
13 and Remy 2011). There is now a general agreement about the global mean
14 sea-level rise (GMSLR) that occurred during the 20th century (see Table 1
15 of Spada and Galassi 2012). However, two related climate issues are still de-
16 bated. The first is the amplitude of the global sea-level acceleration (GSLA)
17 observed during the last centuries and the second is the possible existence
18 of “change points” or “times of inflection” in global reconstructions or in
19 individual tide gauge (TG) records, possibly corresponding to regime shifts
20 of sea-level change. The importance of these issues, both on a regional and
21 on a global perspective, are discussed in the review by Woodworth et al.
22 (2009).

23 In a seminal work, Douglas (1992) estimated the GSLA by averaging
24 the sea-level accelerations obtained from individual records of globally dis-
25 tributed TGs. GSLA is defined as twice the quadratic term in a poly-
26 nomial regression within a limited span of time (henceforth, specific val-
27 ues of GSLA and their uncertainty will be simply denoted by a and Δa ,
28 respectively). The approach of Douglas (1992), similar to that adopted
29 by Douglas (1991) to estimate the secular GMSLR, only provided weak
30 evidence in support to a GSLA, even for the longest period considered
31 (namely $a \pm \Delta a = (0.001 \pm 0.008) \text{ mm/yr}^2$ during 1850–1991). This neatly

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32 contrasted with the significant GSLA predicted to accompany greenhouse
33 warming. The negative result of Douglas (1992) confirmed that of Wood-
34 worth (1990), who limited his attention to European records. No accelera-
35 tion was observed also by Wenzel and Schröter (2010). They reconstructed
36 mean sea level from TGs data (1900–2006) using neural networks although
37 the dataset was then restricted to the period 1950–2006 to prevent the dras-
38 tic reduction of available data during the first half of the century.

39 Recent studies, either based on the “virtual station” stacking method
40 (Jevrejeva et al., 2006, 2008) or on a sea-level reconstruction of long TG
41 records (Church and White, 2006, 2011), unanimously point to the existence
42 of a GSLA. Based on a ~ 300 -years long time series (1700–2002) obtained
43 by combining short and long TG records, Jevrejeva et al. (2008) reported
44 a GSLA of about $a = 0.01 \text{ mm/yr}^2$ (the uncertainty was not quantified),
45 which apparently started at the end of the 18th century. The Empirical Or-
46 thogonal Function (EOF) approach of Church and White (2006), combined
47 with polynomial regression, suggested GSLA of $(0.013 \pm 0.006) \text{ mm/yr}^2$ in
48 the period 1870–2001 and of $(0.008 \pm 0.008) \text{ mm/yr}^2$ when the 20th century
49 only is considered. In the follow-up paper of Church and White (2011),
50 the acceleration $(0.009 \pm 0.003) \text{ mm/yr}^2$ has been proposed for the time
51 period 1880–2009. Sea-level curves previously presented in the literature or
52 obtained in this study are shown in Fig. 1. F1

53 The spread of previous GSLA estimates based on tide gauge (TG)
54 records, summarized in Table 1, is significant. The large energy of decadal T1
55 sea-level fluctuations (Jevrejeva et al., 2006; Chambers et al., 2012; Hous-
56 ton and Dean, 2013), the poor geographical coverage of TGs, the limited
57 number of TGs facing the open seas (hence less affected by coastal pro-

cesses), and the oceans response to regional changes in the pattern of wind stress (Merrifield, 2011; Sturges and Douglas, 2011; Bromirski et al., 2011) are main causes of uncertainty and potential sources of misinterpretation (see also the discussion in Douglas 1992 and Sturges and Hong 2001). As recently evidenced by Gehrels and Woodworth (2013) and by a number of previous studies, the proposed GSLA value is strongly sensitive to the time span of the instrumental record considered and to additional selection criteria based on the quality of the data set. Spurious effects from gappy time series (Wenzel and Schröter, 2010), contaminating tectonic (e.g. Larsen et al. 2003, Olivieri et al. 2013) or anthropogenic factors (Carbognin et al., 2010) act to further complicate the determination of GSLA.

The constant acceleration model for sea-level rise is appealingly simple and constitutes the most obvious generalization of linear models ($a = 0$) extensively employed to estimate GMSLR since the early determination of Gutenberg (1941) (for a review, see Spada and Galassi 2012). However, inspection of sea-level compilations (Gehrels and Woodworth, 2013) and of individual records (see e.g. Bromirski et al. 2011), also reveal short-lived accelerations and abrupt steepness variations. These can be modeled, to a first approximation, as change points (CPs) separating periods of constant rate and/or of constant acceleration. As pointed by Church and White (2006), a CP model including an abrupt slope change at year ~ 1930 , unexpectedly during a period of little volcanic activity, can indeed be invoked as a possible alternative to a constant acceleration model for the time period 1870–2001. Inflections in global and regional compilations of instrumental records at year ~ 1930 have also been proposed by Jevrejeva et al. (2008), Woodworth et al. (2009) and Church and White (2011). Based on proxy

84 and instrumental observations from seven sites, Gehrels and Woodworth
85 (2013) have recently proposed that year 1925 (± 20) could mark the date
86 when sea-level rise started to exceed the long-term Holocene background
87 rate. Inflections or CPs occurring during the 19th century could be more
88 difficult to ascertain in view of the limited amount and sparsity of instru-
89 mental data available for that epoch. However, a major acceleration episode
90 has been evidenced by Jevrejeva et al. (2006) during 1850–1870, though its
91 significance was disputed.

92 Here we provide a new assessment of GSLA based on instrumental (TG)
93 data alone, for the time period 1820–2010. Assuming a constant acceleration
94 model, from a cumulative sea-level curve constructed by TG time series of
95 sufficient length, we obtain GSLA values that are generally consistent with
96 earlier estimates. However, by simple statistical methods, we address in a
97 systematic manner the important role played by non-synchronous CPs at
98 individual TGs in the assessment of the GSLA. Section 2 is devoted to the
99 construction and to the analysis of a global sea-level curve. The results are
100 then discussed in Section 3.

101 2. Results

102 2.1. Building a global sea-level curve

103 In Fig. 1, curves (a) and (b) reproduce the sea-level time series con-
104 structed and studied by Jevrejeva et al. (2006) and by Church and White
105 (2006), respectively. The corresponding GSLA values are given in Table 1.
106 The figure also shows an additional curve (c) that we have built by a global
107 stacking of the 315 Revised Local Reference (RLR) annual time series with

length ≥ 50 yrs, currently available from the Permanent Service for Mean
Sea Level (PSMSL) for the time period 1810–2010 (Woodworth and Player,
2003). It is important to note that we did not apply any low-pass filter to
the selected time series in order to remove multi-decadal fluctuations, as
done by e.g. Jevrejeva et al. (2008). This is motivated by the minimum
length of the time series employed here, which corresponds to the abso-
lute minimum of sea-level record length required to avoid contamination
by low-frequency variations of sea-level (see Fig. 3 of Douglas 1992 and
Jevrejeva et al. 2008). We note, however, that Houston and Dean (2013)
have recently contested this view, proposing that for time series shorter
than 60 years decadal variations significantly affect estimates of underlying
accelerations. Furthermore, we did not attempt to remove *a priori* from
the analysis those TG stations which could be possibly affected by tectonic
movements and particularly those from Japan, which are indeed numerous
(Jevrejeva et al. 2008). As discussed below, the GSLA results obtained here
are largely unaffected by the elimination of stations in tectonically active
areas. The geographical distribution of the 315 stations employed in this
study is shown in Fig. 2a (see also the supplementary kml file). F2a

The stacked sea-level curve (c) in Fig. 1, hereafter referred to as ST
curve, has been obtained by computing the average

$$SL(t_i) = \frac{1}{N(t_i)} \sum_{j=1}^{N(t_i)} (sl_j(t_i) - GIA_j(t_i) - \bar{sl}_j), \quad (1)$$

where $SL(t_i)$ is sea-level at the year $t = t_i$ and $N(t_i)$ is the number of TGs
for which a value of annual mean sea-level is available. The three terms on
the right-hand side of Eq. (1) represent sea-level observed from the j -th
TG at time t_i , the glacial isostatic adjustment (GIA) correction for the j -th

132 TG, and the average sea-level observed during the whole time span during
 133 which the TG has been operating, respectively (note that the subtraction of
 134 $\overline{\text{sl}}_j$ has no influence on the assessment of GSLA). The range of uncertainty
 135 of the global sea-level curve (1) is evaluated by twice the standard deviation
 136 $\text{SLD}(t_i)$ around $\text{SL}(t_i)$. Since the computed value of $\text{SLD}(t_i)$ largely
 137 exceeds the error on the annual mean from individual stations, which can be
 138 estimated at the level of 0.5 mm (Fabio Raichich, personal communication,
 139 2013), this latter is not taken into account in the assessment of the uncertainty
 140 associated with curve ST and with other global or individual time
 141 series considered in the following. In building the stack (1) only years with
 142 $N(t_i) \geq 2$ are considered. The GIA correction has been performed adopting
 143 model ICE-5G(VM2) of Peltier (2004) by means of an improved version
 144 of program SELEN (Spada et al., 2012), originally proposed by Spada and
 145 Stocchi (2007). Possible uncertainties on $\text{GIA}_j(t_i)$ have not been taken into
 146 account.

147 The stacking technique is commonly employed in seismic data processing
 148 in order to increase the signal-to-noise ratio and to enhance the coherency
 149 of time series (see e.g. Gilbert and Dziewonski 1975). We are aware that the
 150 conventional un-weighted stacking (Eq. 1) is not always satisfactory and
 151 better results can be obtained using more sophisticated averaging techniques
 152 (Liu et al., 2009). However, our elementary approach to the construction of
 153 a global sea-level curve is motivated, *a-posteriori*, by the consistency of the
 154 GSLA obtained in this way with previous independently derived estimates.

155 Using bootstrapping (Efron and Tibshirani, 1986), we have determined
 156 the best-fitting quadratic polynomial for curve ST, which reads:

$$\text{SL}(t) = (0.0049 \pm 0.0012) t^2 + p_1(t), \quad (2)$$

157 where $SL(t)$ is expressed in mm and t in years, and $p_1(t)$ is a degree one poly-
 158 nomial whose coefficients are not of concern here. The uncertainty on the
 159 quadratic term corresponds to the rms of the distribution of 5,000 quadratic
 160 terms obtained by synthetic curves where $SL(t_i)$ is taken randomly from a
 161 Gaussian deviate with standard deviation $2SLD(t_i)$. The GSLA implied in
 162 (2), i.e. twice the quadratic term, is (0.0098 ± 0.0023) mm/yr², where the
 163 uncertainty corresponds to the 95% confidence interval, as in Church and
 164 White (2006) (here and in the following, we round off to two significant
 165 figures in the GSLA uncertainty when the leading digit is 1 or 2, see e.g.
 166 Taylor 1997).

167 Despite the crude averaging implied in Eq. (1), result (2) is generally
 168 coherent with previous findings and constitutes an independent confirma-
 169 tion of the existence of a GSLA. In particular, our estimate well matches
 170 the value $a \sim 0.01$ mm/yr² proposed by Jevrejeva et al. (2008). Since an
 171 estimate of the associated uncertainty is not provided by Jevrejeva et al.
 172 (2008), we have applied the bootstrapping procedure to their original data.
 173 This gives $\Delta a = 0.002$ mm/yr², in close agreement with our estimate above
 174 based on stacking. This result, however, should be taken cautiously in view
 175 of the significantly longer record considered by Jevrejeva et al. (2008) and
 176 the larger number of TGs utilized (1023 stations versus the 315 employed
 177 here). While the agreement of our result with Church and White (2006)
 178 (i.e. $a = 0.013 \pm 0.006$ mm/yr²) is satisfactory, we note that our GSLA es-
 179 timate above turns out to be more precise (the fractional uncertainty is
 180 $\Delta a/a \sim 20\%$) than in Church and White (2006) (fractional uncertainty
 181 $\sim 50\%$). Since the two methods and the two TG sets employed differ, the
 182 origin of this discrepancy is difficult to assess. This, of course, also holds for

the overall accuracy of our estimate. It is expected that the record length plays a major role in increasing the uncertainty of the assessment. We could verify that when the ST curve is restricted to the same time period (1870–2001) considered by Church and White (2006), the fractional uncertainty increases to $\sim 50\%$.

To evaluate the impact of the number of TGs employed in the GSLA assessment and the related selection criteria, in Fig. 1 we show two further synthetic sea-level curves obtained by Eq. (1). The first (curve d) has been constructed using the global set of 23 TGs considered in the GMSLR assessment of Douglas (1997) henceforth referred to as D97 set, while the second (curve e) includes the 22 TGs recently employed in the study by Spada and Galassi 2012, herein referred to as SG01 TG set. These two global sets, which are partly overlapping, have been determined imposing specific constraints to the length and to the quality of the TG time series (D97) and, in addition, requiring that the GMSLR estimate is essentially independent upon the GIA correction adopted (SG01). We remark that TG stations which could be possibly affected by tectonic movements are expunged *a priori* from these two sets. The bootstrapping procedure provides, for the two sets, consistent GSLA estimates, namely (0.012 ± 0.002) mm/yr² and (0.013 ± 0.002) mm/yr², respectively. These agree with the results based on the ST curve and with previous estimates in Table 1. This finding supports the idea that, similarly to GMSLR, GSLA can be detected even using a limited number of TGs, provided that their spatial coverage is sufficient and rigorous selection criteria are imposed. This is only apparently in contradiction with the seminal work of Douglas (1992), who effectively imposed these criteria. His negative result with respect to the existence of a GSLA was

209 likely due to the shorter time series compared to those available nowadays.

210 *2.2. Analysis of the sea-level curve ST*

211 In Fig. 3, curve ST is studied more in detail. Red dotted lines above and F3
212 below the curve correspond to one standard deviation $SLD(t_i)$. Since the
213 number of TGs operating every year $N(t_i)$ varies considerably with t_i (the
214 dependence is displayed in the bottom part of Fig. 3), $SLD(t_i)$ is markedly
215 time dependent. This feature, which also characterizes the reconstructions
216 of Jevrejeva et al. (2008) and reflects the non-stationarity of the time series,
217 has an important role in the assessment of the best-fitting curve and of the
218 uncertainty on the corresponding GSLA.

219 To better scrutinize the nature of the non-linear trend shown by the ST
220 curve in Fig. 3, we have compared the results of the quadratic regression,
221 expressed by Eq. (2), with those obtained by a linear and a bi-linear re-
222 gression, respectively. The adoption of a bi-linear model is motivated by
223 previous studies (e.g. Church and White 2006), which have evidenced the
224 existence of CPs or “inflections” in global sea-level curves, corresponding
225 to abrupt slope variations, hence to short-lived sea-level accelerations. A
226 review of the literature supporting the existence of CPs is given by Gehrels
227 and Woodworth (2013). Furthermore, CPs are also suggested by modeled
228 scenarios of future sea-level rise (Spada et al., 2013), and particularly by
229 the sea-level component expected from terrestrial ice melt, showing abrupt
230 changes of the sea-level trend in response to episodes of enhanced mass loss
231 in Greenland. For the sake of parsimony, we have not attempted to in-
232 troduce more sophisticated multi-linear regression methods, which appears
233 unmotivated in view of the large errors generally affecting the construction

234 of a global sea-level curve. However, the use of multi-linear models could
235 be appropriate when regional secular sea-level records are considered, as in
236 the case of the Pacific coast of North America (Bromirski et al., 2011).

237 Here, bi-linear regression has been performed adopting methods em-
238 ployed in econometrics to detect structural changes, i.e. variations of the
239 statistical parameters of non-stationary time series such as ST including,
240 in particular, changes in the rate of variation. The Chow statistics allows
241 for the detection of a CP at a given time (Chow, 1960). In this testing pro-
242 cedure, the time series is split into two sub-periods, and for each of them
243 a linear regression is performed. Continuity is not imposed at the time of
244 occurrence of a CP. The misfit obtained for such bi-linear model is then
245 compared, by means of a Fisher F-test (e.g. Winer 1962), with the one
246 obtained by a linear model for the whole time series. We have implemented
247 the recipe by Hansen (2001), based on an idea of Quandt (1960), which
248 overcomes the limitation caused by the need for the break date to be known
249 *a priori* and introduces a methodology for determining a structural change
250 whose timing is unknown.

251 Analysis of the ST curve shows that the bi-linear regression significantly
252 improves the fit (at the confidence level $\alpha = 95\%$) with respect to a linear
253 or a quadratic model. The structural CP, which corresponds to the largest
254 value of the Chow statistics, is found within the time interval 1835–1840.
255 This is relatively close to the sea-level acceleration visually evidenced by
256 Jevrejeva et al. (2006) in the period 1850–1870, which appears to be the
257 largest acceleration visible in the sea-level reconstruction during the last
258 200 years (see their Figure 5). Since the dataset employed and the methods
259 of analysis differ, we tentatively suggest that the CP we have detected effec-

260 tively corresponds to the acceleration episode described by Jevrejeva et al.
 261 (2006). This would constitute, indirectly, a validation of the automated CP
 262 search method adopted here. Though the misfit reduction obtained for the
 263 bi-linear model is indeed statistically significant (at the 95% significance
 264 level) compared to a linear or quadratic regression, the global nature of
 265 this CP is dubious. The reason is that only data from six PSMSL sta-
 266 tions clustered in Europe contribute to the ST curve in the lapse of time
 267 between 1830 and 1849 (namely, Brest (F), Swinoujscie (PL), Sheerness
 268 (GB), Cuxhaven 2 (D), Wismar 2 (D), and Maassluis (NL), see Fig. 2b). F2b
 269 By similar arguments, Jevrejeva et al. (2006) have pointed to the dubious
 270 significance of the acceleration episode, since only five stations, facing the
 271 North Atlantic and the Baltic, were in operation. Visual inspection of the
 272 six records above corroborate the hypothesis of a CP in the earliest time
 273 series (Brest and Swinoujscie), which is also confirmed by a separate anal-
 274 ysis. The commencement of the remaining four records around year 1850,
 275 followed by a marked and coherent linear sea-level rise, acts to strengthen
 276 the 1835–1840 structural change.

277 To avoid any bias resulting from a poor spatial coverage of the stacked
 278 time series, hereinafter we will consider the second branch (referred to as
 279 ST2) of ST, pertaining to the time period 1840–2010; this curve is shown in F4
 280 Fig. 4. The number of RLR records that build ST2 progressively increases
 281 to ~ 300 until ~ 1960 , and decreases to ~ 200 by the year 2010. As we
 282 have verified, a sufficient spatial coverage is ensured for the TGs used to
 283 construct curve ST2, with no clusterings at continental or regional scales
 284 scale during the whole time span. Analysis of curve ST2 reveals that the

285 quadratic regression

$$\text{SL}(t) = (0.0021 \pm 0.0012) t^2 + q_1(t), \quad (3)$$

286 where SL is expressed in mm and t in years and $q_1(t)$ is a linear polynomial,
287 improves the fit ($\alpha = 95\%$) with respect to linear and bi-linear models lim-
288 ited to the same period. Eq. (3) implies an acceleration (0.0042 ± 0.0024)
289 mm/yr², which confirms the existence of a GSLA and points, in particular,
290 to the absence of significant CPs during 1840–2011. To test the robustness
291 of these results against the number of TGs used for a given year, we have
292 stacked time series with length ≥ 60 and ≥ 75 years (the number of used
293 TGs reduces from 315 to 225 and 143, respectively). These computations
294 confirm the existence of the GSLA and the absence of significant CPs, show-
295 ing that Eq. (1) is not introducing artifacts when there is a change in the
296 TGs available at a given time. This has also been confirmed by further
297 computations, in which following Jevrejeva et al. (2006) we have performed
298 the stacking on the rates of each individual time series. The derivatives
299 have been numerically implemented using a two-points, two-sided formula.
300 The resulting sea-level curves essentially reproduce the time-derivatives of
301 our curves ST and ST2, thus showing that no artifacts are introduced when
302 a change in the number of TGs available occurs.

303 Although the ST2 curve is best-fitted (95%) by a parabola (see Eq. 3),
304 it is of interest to determine the best-fitting bi-linear model. When this
305 is done, the CP is found for year ~ 1940 , relatively close to the inflection
306 evidenced by Jevrejeva et al. (2008), Woodworth et al. (2009) and Church
307 and White (2011) for year ~ 1930 . For consistency with our statistical ap-
308 proach, this can only be classified as a “weak” CP with low-significance,

309 since the best performing model is, for curve ST2, the quadratic one given
 310 by Eq. (3). As pointed by Rahmstorf (2007), the variation of the trend of
 311 sea-level rise that occurred in ~ 1940 corresponded to a major variation of
 312 the global temperatures. We observe that the GSLA value implied in our
 313 estimate (3) turns out to be ~ 3 times smaller than the previous estimate
 314 $(0.013 \pm 0.006 \text{ mm/yr}^2)$ by Church and White (2006), which covers a compa-
 315 rable time span, but was obtained by distinct selection criteria and methods
 316 of analysis (see Table 1). The precisions of the two estimates, measured by
 317 their fractional uncertainty ($\sim 50\%$), are comparable.

318 *2.3. Interpreting the sea-level curve*

319 Averaged expressions like (2) and (3), based on the stacking (1), are
 320 appealing, since they are supposed to capture the actual ocean behavior in
 321 an apparently simple fashion. However, regardless the averaging method-
 322 ology adopted, these approaches tend to hide the mechanisms that control
 323 the local sea level change recorded in single time series. For instance, the
 324 effective source of the quadratic trend itself remains obscure, until the indi-
 325 vidual components of the stacking are scrutinized or the forcing mechanism
 326 is identified. The quadratic growth of curve ST2 does not necessarily imply
 327 a similar behavior for all the time series that compose the stack, although
 328 one could intuitively expect that a dominance of quadratic time series would
 329 be ultimately responsible for the observed GSLA.

330 To address the issues above, we have classified the 315 TGs that form
 331 curves ST and ST2 according to the regression models that best fit each
 332 time series that contributes to the stacking. The best fitting models have
 333 been determined by ordinary least squares, since serial correlation has been

334 shown by Baki Iz et al. (2012) not to affect the estimates of the trends,
 335 while the impact on the uncertainty is minor. With TG-L, TG-Q and
 336 TG-B we indicate time series subsets for which the best-fitting statistical
 337 model is linear, quadratic and bi-linear, respectively. The performances of
 338 these models have been compared analyzing the variances of the residuals
 339 by means of a F-test (with $\alpha = 95\%$). The most populated subset is TG-
 340 L, which contains 237 time series (75% of the total); most of them (90%)
 341 show a positive sea-level trend. Subset TG-Q contains only 47 time series
 342 (15% of the total). With the exception of three sites, all the time series
 343 belonging to set TG-Q show a positive trend (i.e. the linear term of the
 344 quadratic model) and most of them (75%) are characterized by a positive
 345 quadratic term (i.e. $a > 0$). Finally, subset TG-B only contains 31 time
 346 series ($\sim 10\%$ of the total). The CPs of the TG-B time series are marked
 347 by vertical bars in Fig. 4, where red and blue colors imply an increase and
 348 a decrease of the sea-level trend across the CP, respectively. CPs show a
 349 complex temporal distribution, but some patterns emerge. They appear
 350 only sporadically before ~ 1960 while they are more frequent and energetic
 351 afterwards and particularly during the last four decades. Furthermore, red
 352 CPs dominate the blue ones in terms of amplitude and frequency (24 out
 353 of the 31 CPs detected are red).

354 It is worth to recall that the 315 TGs employed to construct curve ST2
 355 were selected only according to the record length criterion. When the anal-
 356 ysis performed on the records contributing to ST2 is extended to the D97
 357 (Douglas, 1992) and SG01 (Spada and Galassi, 2012) time series, for which
 358 additional selection criteria have been applied, similar results are found.
 359 Namely, most of the time series are best-fitted by a linear polynomial (77%

and 81% of the total number of time series for sets D97 and SG01, respectively). The few remaining are almost equally partitioned in two sets, fitted by quadratic and bi-linear models, respectively. This confirms that sets D97 and SG01 are effectively representative of the global set of TGs, also with respect to the style of the statistical models that best-fit their components.

The spatial distribution of TGs belonging to the TG-L, TG-Q and TG-B subsets is shown in Fig. 5. The dominance of positive sea-level trends (red dots) for TG-L stations is apparent in Fig 5a. Negative trends (blue dots) are mainly clustered regionally. These are observed along the North and the South American West coast, where they can be considered, at least partly, as the result of active tectonics along transcurrent and collisional boundaries in these regions. Negative sea-level trends along the Pacific coasts of North America since ~ 1980 have been recently attributed to a steric response to wind stress, and interpreted as indications of an imminent sea-level acceleration (Bromirski et al., 2011). Wind stress has been also recognized as the source of large sea-level drops in the eastern North Pacific and North Atlantic coasts between the late 1800s and the early 1900s (Sturges and Douglas, 2011). Negative rates of sea-level change observed in northern Europe and particularly along the coasts of the Baltic Sea can be associated with the ongoing post-glacial rebound in response to the melting of the late-Pleistocene ice sheets (see e.g. Spada and Galassi 2012).

Because of their relatively small number compared to TG-L, spatial patterns in the distribution of the TG-Q (Fig. 5b) and of the TG-B TGs (5c) cannot be easily identified. This would suggest that the positive acceleration expected from a stacking of the TG-Q time series does not have a regional origin. It is remarkable that the Japanese TGs show trends of all

the three kinds so far discussed. This is likely to reflect the complex tectonic setting of this region (Aubrey and Emery, 1986), which makes the interpretation of the TG signals particularly difficult (see e.g. the discussion in Spada and Galassi 2012). The sea-level time-series for San Francisco falls in the TG-B category (5c). For this record, our analysis indicates a CP for year ~ 1890 . According to Bromirski et al. (2011), who limited their attention to the last century, two major discontinuities in the rate of sea-level rise can be evidenced for San Francisco at times ~ 1930 and ~ 1980 , which are also visible in the San Diego and Seattle records. In other approaches, based on the smooth Intrinsic Mode Functions (Breaker and Ruzmaikin, 2013), abrupt CPs could not be resolved for the San Francisco record, although their existence is strongly suggested by a visual inspection of the full time series, after the application of a running average filter. Finally, we note that according to our analysis, none of the TGs located in the Pacific area belongs to the TG-B subset. Indeed, application of the “virtual station method” to TG records from this region reveals complex regional patterns that could hardly be consistent with a single-CP regression model (Webb and Kench, 2010).

Using Eq. (1), the time series belonging to the three subsets TG-L, TG-Q and TG-B have been stacked and the resulting global curves have been analyzed in order to determine the best fitting statistical model. This aims F6 to check how different styles contribute to GSLA. The results are shown in Fig. 6. As expected, the stacked TG-L time series are best fitted ($\alpha = 95\%$), by a linear model. Its regression coefficient corresponds to a rate of sea-level rise of (0.94 ± 0.11) mm/yr. Similarly, the stack obtained using TG-Q data are best fitted by a quadratic polynomial that implies a sea-level ac-

412 celeration (0.003 ± 0.005) mm/yr². The stacked TG–B time series, however,
 413 demand a quadratic model as well, corresponding to a sea–level acceleration
 414 (0.012 ± 0.006) mm/yr². For the stacking resulting from the TG–B set, the
 415 rejection of a bi–linear model can be interpreted as the cumulative effect
 416 of the time sequence of essentially coherent change points that characterize
 417 the TG–B time series (their timing and amplitude are shown in Fig. 4).
 418 The resulting stacked curve is best–fitted by a parabola characterized by
 419 a positive acceleration, matching the envelope of several time series having
 420 the shape of linear segments separated by non simultaneous CPs. This is
 421 not totally unexpected and was remarked (but not made quantitative) by
 422 Gehrels and Woodworth (2013) when discussing the local contribution to
 423 global instrumental sea–level curves.

424 **3. Discussion and conclusions**

425 Un–weighted stacking of the longest RLR annual TG time series pro-
 426 duces a synthetic global sea–level curve (ST), which shows several features.
 427 First, ST shows a statistically significant and positive CP, implying a sud-
 428 den increase in slope, within the time period 1835–1840. Second, branch
 429 ST2 of curve ST, which encompasses the time period 1840–2010, is best
 430 fitted by a quadratic polynomial ($\alpha = 95\%$). This confirms previous results
 431 about the existence of a GSLA for the period 1840–2010 (Jevrejeva et al.,
 432 2006; Church and White, 2006). According to our estimates, the GSLA is
 433 (0.0042 ± 0.0024) mm/yr². The projection of Eq. (2) to year 2100 suggests
 434 a sea–level rise of about 16 cm relative to 1990, at the lower boundary of
 435 the IPCC projection for 2100 of the observed sea–level rise from the 20th
 436 century, which is in the range of 19–58 cm (Meehl et al., 2007). This es-

437 timate would increase to 22 cm using Eq. (2), which represents the best
438 fitting parabola for curve ST (see the supplementary Fig. S1). F1S

439 The determination of the starting point of the present sea-level rate
440 and acceleration is one of the challenges of current studies since it could
441 unveil correlation with anthropogenic factors or global climate change. The
442 most recent and comprehensive study by Gehrels and Woodworth (2013)
443 has proposed the existence of a possible sea-level inflection at the year
444 1925 ± 20 . This time window includes the early result by Woodworth (1990)
445 who proposed year 1930 for the inflection, which was subsequently confirmed
446 by Church and White (2006) and Woodworth et al. (2009). Since the sea-
447 level curve ST2 is best-fitted by a quadratic model, our statistical analysis
448 does not support the existence of a CP. However, we have verified that
449 among all the possible bi-linear models for ST2, the residues are minimized
450 when a CP at year ~ 1940 is allowed, which could be assimilated to the one
451 evidenced in the previous literature.

452 Results by Jevrejeva et al. (2006) suggest a major change in the rate of
453 sea-level change during the period 1850–1870, which probably marks the
454 start of present acceleration. Our curve ST evidences a CP between 1835
455 and 1840 which could be interpreted as a new CP not observed before. It is
456 possible, however, that here we are observing the same short-term acceler-
457 ation detected Jevrejeva et al. (2006). The non exact temporal coincidence
458 of the two episodes could be justified by the different sets of TGs employed
459 and the different approaches. Furthermore, while in Jevrejeva et al. (2006)
460 the acceleration episode has been identified visually, here we have used an
461 automatic search strategy. Our analysis shows that it is impossible to ascer-
462 tain the global origin of this CP (or of these CPs), since the few operating

463 TGs in that period were located in Northern Europe. Similar conclusions
464 have been drawn by Jevrejeva et al. (2006).

465 The results presented in this work are suggesting a re-evaluation of the
466 same meaning of GSLA. In fact, though the best-fitting model for curve
467 ST2 is indeed quadratic in the time period 1840–2010, our analysis has
468 shown that most of the components follow a linear model. The number of
469 effectively quadratic time series (TG–Q) is limited to $\sim 15\%$ of the total
470 of 315 RLR time series considered in this study. Although the number
471 of bi-linear ones (TG–B) is even smaller ($\sim 10\%$), the time sequence of
472 CPs provides the stack an upward curvature that enhances the effect of the
473 TG–Q time series and coherently emerges from the averaging. One of the
474 reasons the acceleration only emerges in a limited number of sites ($\sim 25\%$)
475 is that long and very long period oscillations dominate the signal while
476 nodal points are scarce and unlikely to coincide with all or just some of the
477 selected observation points.

478 The findings above, obtained by the application of a modified Chow test
479 (Hansen, 2001), have two important consequences. *i)* when dealing with
480 GSLA, the attribute *global* should be used cautiously, since the vast major-
481 ity of the TG time series used to construct the ST curve are effectively *not*
482 showing any significant acceleration ($\alpha = 95\%$). Indeed, from an analysis of
483 the distribution of the TG–Q instruments (see Fig. 5), we have found that
484 these are often surrounded by sites that do not show any significant accel-
485 eration (see Supplementary Material). The *global* nature of the GSLA only
486 stems from the lack of any apparent regional clustering in the spatial dis-
487 tributions of the TG–Q and TG–B gauges (see Fig. 5), which ultimately
488 determine the parabolic shape of the cumulative curve. *ii)* intermittent and

489 non-synchronous CPs occurring at individual TG since ~ 1880 have an
490 important role in determining an *average* sea-level acceleration on a cen-
491 tury time scale. The relevance of “short-term accelerations” enlightened in
492 several previous studies (see Gehrels and Woodworth 2013 and references
493 therein) is therefore confirmed in this study. Here, the problem has been
494 put in a quantitative perspective using statistical methods borrowed from
495 econometrics.

496 **4. Acknowledgments**

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502 manent Service for Mean Sea Level) archive on August 1st, 2012 ([http://](http://www.psmsl.org/data/obtaining/)
503 [www.psmsl.org /data /obtaining/](http://www.psmsl.org/data/obtaining/)). All figures have been drawn using the
504 Generic Mapping Tools (GMT) (Wessel and Smith, 1998).

Table 1: GSLLA estimates based on global analyses of instrumental records, for which a quadratic term in a polynomial regression is evaluated.

Author(s)	year	$a \pm \Delta a$ (mm/yr ²)	Period (year–year)	Input dataset	Methods
Douglas	1992	-0.011 ± 0.012	1905–1985	23 TGS	Averaged group acceleration
”	”	0.001 ± 0.008	1850–1991	37 TGS	”
Church and White	2006	0.013 ± 0.006	1870–2001	PSMSL TGS	EOFs
”	”	0.008 ± 0.008	20 th century	”	”
Jevrejeva et al.	2008	~ 0.01	1700–2002	1023 RLR PSMSL TGS	Virtual station method
Church and White	2011	0.009 ± 0.003	1880–2009	PSMSL TGS and NASA/CNES altimetry	EOFs
This study	2013	0.0098 ± 0.0023	1820–2010	315 PSMSL RLR TGS	Stacking (ST curve)
This study	2013	0.0042 ± 0.0024	1840–2010	”	Stacking (ST2 curve)

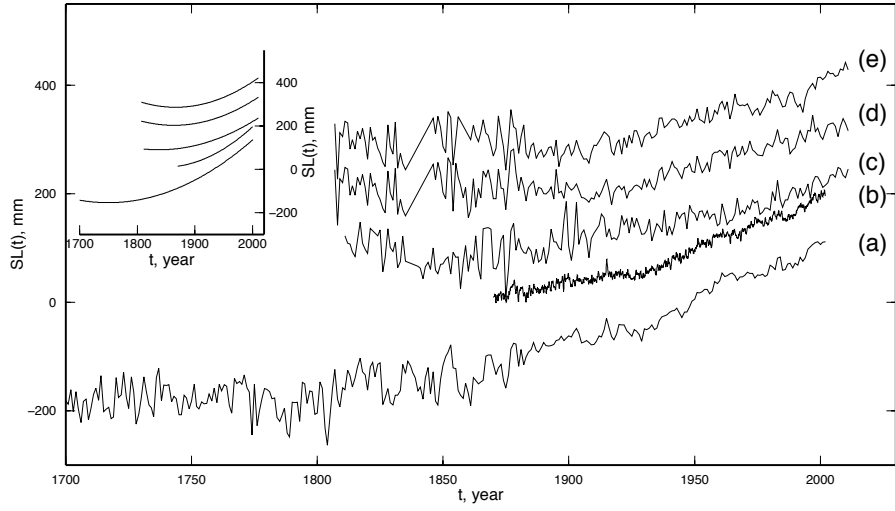


Figure 1: Various sea-level curves relevant to this work (the curves are shifted by an arbitrary amount for to facilitate visualization). Curves (a) and (b) show the reconstructions by Jevrejeva et al. (2006) (the standard errors are not reproduced from the original work) and Church and White (2006), respectively. Curve (c) is the ST time series obtained in this work by the stacking of RLR TG observations. Curves (d) and (e) result from the stacking of the TGs selected by Douglas (1992) and by Spada and Galassi (2012). The best-fitting quadratic polynomials to curves (a–e) are shown in the inset, while numerical values of the corresponding accelerations are given in Table 1.

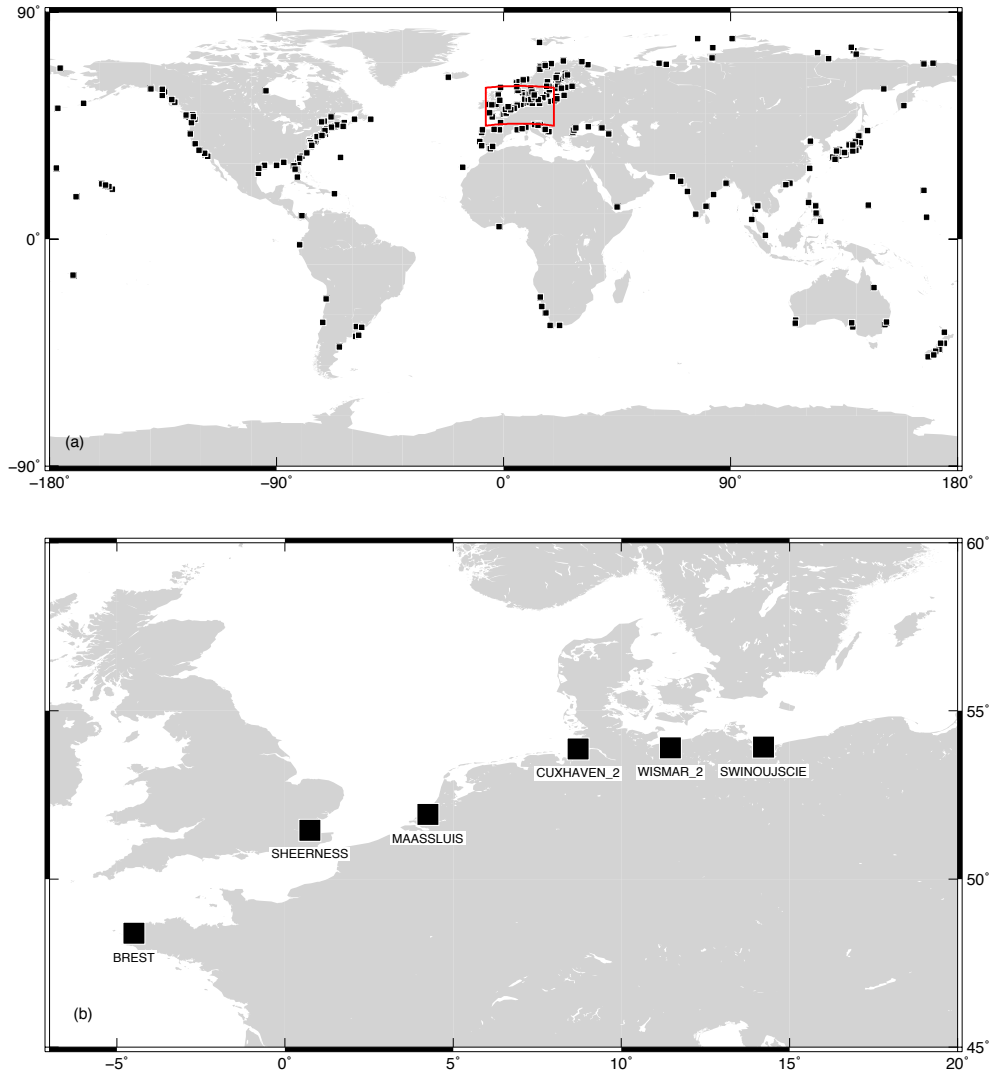


Figure 2: (a) Geographical distribution of the 315 RLR TGs for which > 50 years of data are available within the period 1820–2010, which build the ST curve. The region in the inset is enlarged in (b), and shows the location of the six North European TG rime series available in the period 1830–1849.

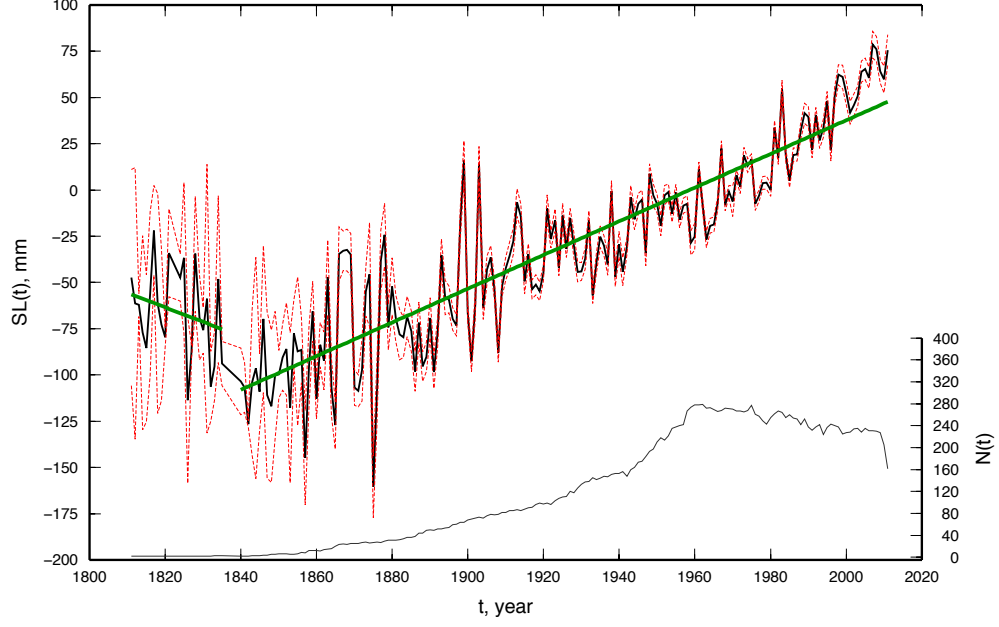


Figure 3: The stacked curve ST (black), obtained from Eq. (1) for the period 1810–2010 and the range of uncertainty corresponding to $SLD(t_i)$ (red). The green line represents the best fitting bi-linear model for ST , showing a CP for year 1835–1840. The regression coefficient rises from (-1 ± 3) mm/yr before the CP to (0.91 ± 0.05) mm/yr after the CP. The plot at the bottom shows $N(t)$, the number of time series available in the stacking at a given epoch t .

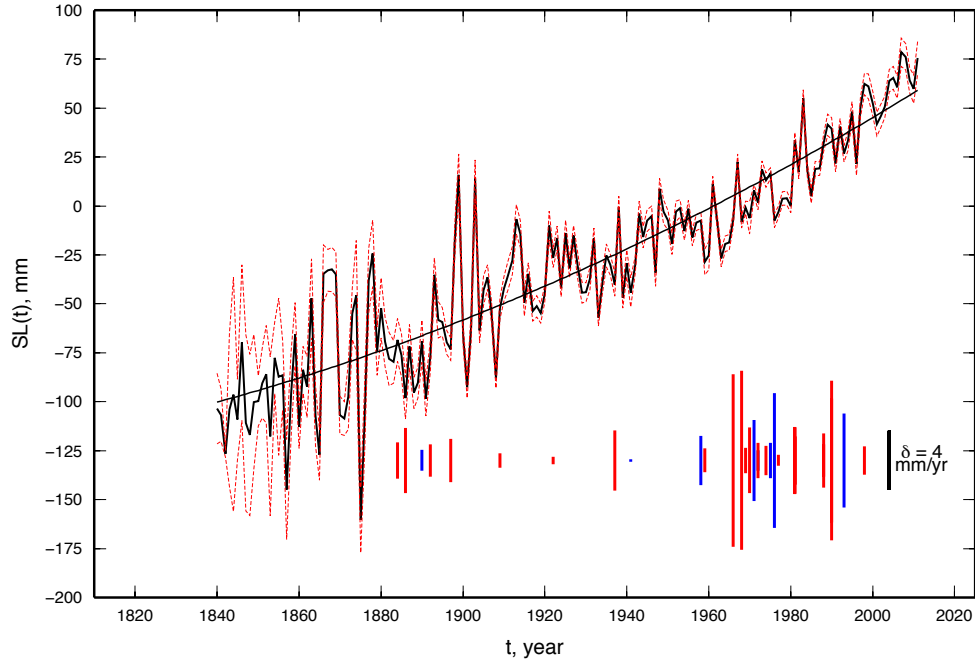


Figure 4: The same as in Fig. 3, but for curve ST2, obtained from Eq. (1) for the period 1840–2010. The black curve shows the best fitting quadratic polynomial. Vertical bars at the bottom of figure show the sequence of CPs found for each of the time series in the TG–B set. Red and blue segments indicate CPs for which the variation in the rate of sea–level change, denoted by δ , is positive and negative, respectively.

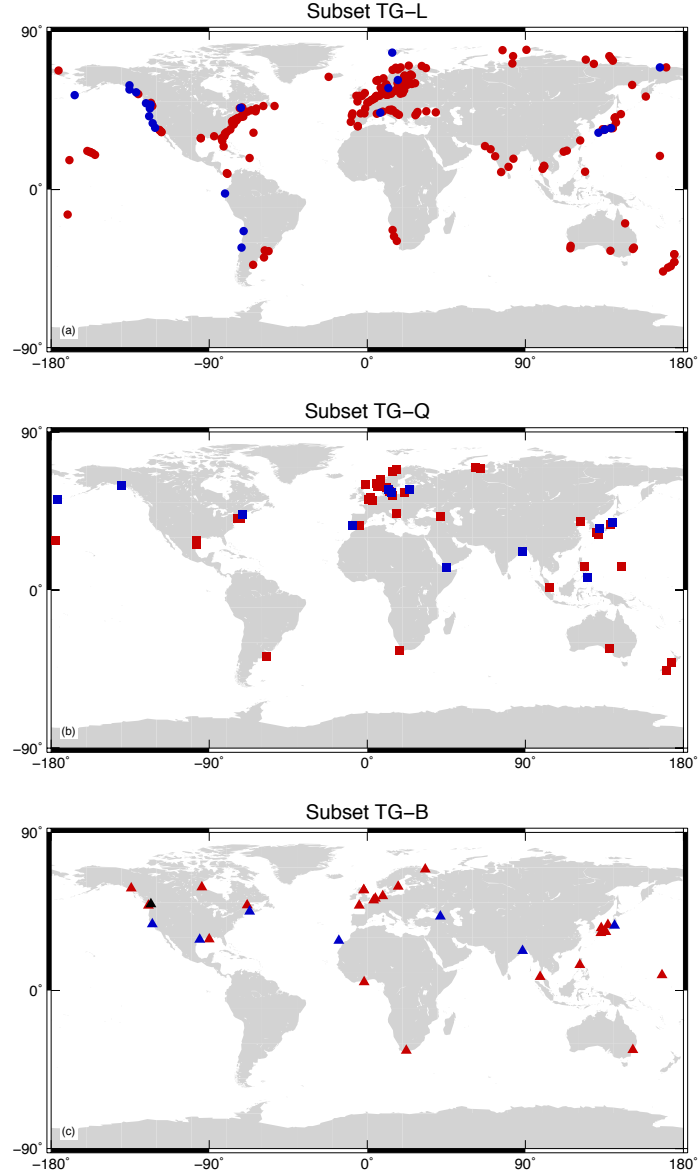


Figure 5: Locations of TGs according to the best-fitting model. (a): Subset TG-L (red and blue symbols denote positive and negative trends, respectively), (b): TG-Q (the red color indicates a positive quadratic term, the blue a negative one), (c): TG-B (red and blue colors indicate positive and negative values of δ (see Fig. 4)).

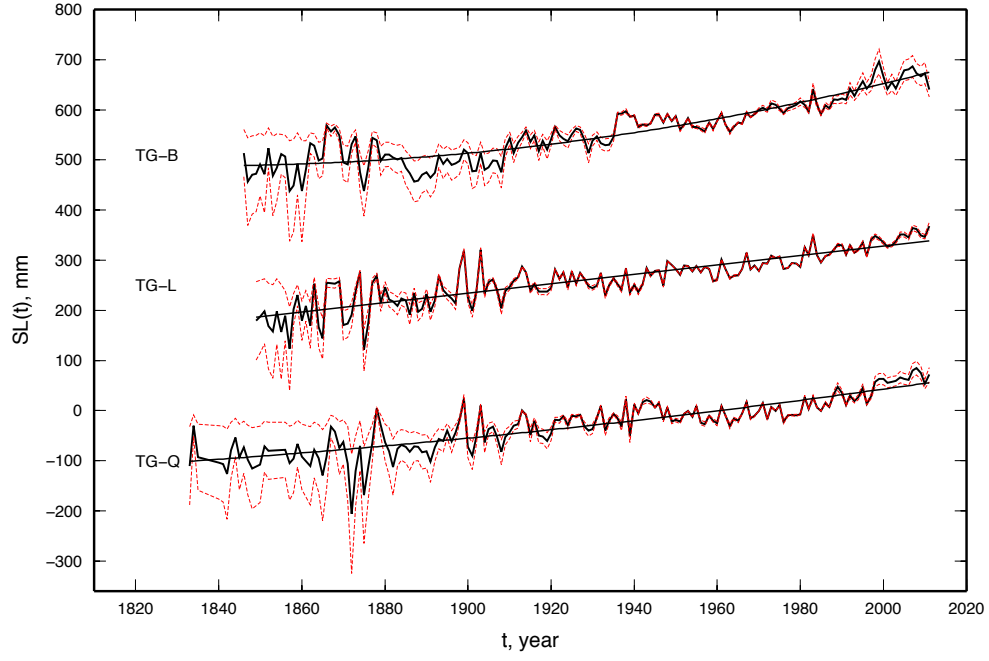


Figure 6: Stacking obtained for the time series belonging to the TG-B (top), TG-L (middle) and the TG-Q (bottom) subsets. For the three subsets, the best fitting models are quadratic, linear and quadratic, respectively. Red dashed curves mark the 1σ interval for the stacks.